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MONETARY TRANSMISSION IN THREE CENTRAL EUROPEAN ECONOMIES: EVIDENCE FROM TIME-VARYING COEFFICIENT VECTOR AUTOREGRESSIONS

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Highlights

- We study the transmission of monetary policy to macroeconomic variables with structural time-varying coefficient vector autoregressions in the Czech Republic, Hungary and Poland, in comparison with that in the euro area. These three countries have experienced changes in monetary policy regimes and went through substantial structural changes, which call for the use of a time-varying parameter analysis. Our results indicate that the impact on output of a monetary shock changed over time. At the point of the last observation of our sample, the fourth quarter of 2011, among the three countries, monetary policy was most powerful in Poland and not much less strong than the transmission in the euro area. We discuss various factors that can contribute to differences in monetary transmission, such as financial structure, labour market rigidities, industry composition, exchange rate regime, credibility of monetary policy and trade openness.

Keywords: monetary transmission; time-varying coefficient vector autoregressions; Kalman-filter.
JEL codes: C32, E50

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1 Introduction

The monetary transmission mechanism describes the effects of monetary policy on key macroeconomic and financial variables. Analysis of the monetary transmission mechanism has a central role in macroeconomic policy research and is crucial for the conduct of monetary policy. A good understanding of the transmission mechanism is especially important for the implementation of inflation targeting. As the transmission mechanism changes, the reaction function of monetary policy should also change, even if the central bank's preferences are unchanged.

The monetary policy transmission mechanism also has implications for euro adoption and proper functioning within a monetary union. The relevance of the monetary transmission mechanism from the perspective of euro adoption is that when the effect of domestic monetary policy on the economy is large, is substantially different from the effects observed in the euro area and business cycles are not sufficiently synchronised, then the cost of losing monetary-policy independence might be significant. In the opposite case, the loss is less important. On the other hand, similarity of the transmission mechanism across member states of a currency area is important. Ideally, monetary policy should have the same effect on all member states, causing them to share equally the burden of adjustment after a monetary contraction or the advantages of a monetary easing. Analysis of the transmission mechanism for countries currently outside the monetary union can give insights into euro-adoption considerations, but has less relevance for the question of proper functioning within a monetary union, because the transmission mechanism may change after entry into the monetary union.

The Monetary Transmission Network of the European System of Central Banks (ESCB) analysed in detail the transmission mechanism in the current euro-area member countries (Angeloni *et al*, 2003). A large amount of research has also been conducted for the central European member states of the EU; see Égert and MacDonald (2009) for an extensive survey. However, studying the central European countries with standard techniques raises an elementary problem that is related to the possibility of time-varying parameters, for three reasons.

First, both common sense and the critique of Lucas (1976) suggest that changes in monetary regimes are likely to affect the transmission of monetary policy. Lucas (1976) concluded that any change in economic policy will systematically alter the structure of econometric models. For example, a shift from exchange-rate targeting to inflation targeting may weaken exchange rate pass-through, because an exchange rate change can be more persistent in the former than in the latter regime. During the last two decades a number of central European countries, such as the Czech Republic, Hungary and Poland, made their exchange rate regimes more flexible and changed their ways of conducting monetary policy, ultimately arriving at an inflation-targeting framework. Regime changes call into question the usefulness of studying the available sample period of these countries with fixed parameter models, because the period since the last change typically does not provide a sufficient number of observations for estimation. When the sample available since the last regime change seems sufficient to conduct a fixed-parameter analysis, such as in the case of the Czech Republic, then time-varying parameter analysis allows the study of the change in monetary transmission due to shocks, such as the 2007-08 global financial and economic crisis. Time-varying parameter analysis also allows consideration of the issues described in the next two paragraphs.

Second, these countries have undergone substantial structural changes since their transition from the socialist economic system and integration into the EU, which affected, among others factors, trade and financial linkages with western European countries, market regulations and institutional structures. These changes might have affected the parameters of response function of monetary policy even if the monetary regime was unchanged, such as in the Czech Republic during the last decade.

Third, the parameters of the linear models that are typically used[1], can change even if the underlying structural model has constant parameters, provided that the underlying structural model is nonlinear. This is a major reason for Granger (2008) advocating the use of time-varying parameter techniques.

Despite the above-mentioned motivations, we aware of only one (not-yet-published) paper, Franta, Horváth and Rusnák (2011), which adopted time-varying coefficient (TVC) techniques for studying the transmission mechanism of a central European country, the Czech Republic.

Yet possible time variation in the transmission mechanism is not only a central European issue. There is an intensive debate about the possible changes in monetary transmission of the USA and the euro area based on time-varying coefficient analysis. Canova and Gambetti (2006) question whether US inflation is the result of "bad luck or bad policy," that is, whether the bad inflation outcome of the early 1980s and the good inflation outcome since the late 1980s are due to "luck" (the decline in the variance of the shocks) or to "policy" (changes in the way monetary policy is conducted). They found more evidence in favour of the bad luck hypothesis, while Cogley and Sargent (2005) support the bad policy view. The empirical results of Primiceri (2005) are also more supportive of the bad luck view, finding that the role played by exogenous non-policy shocks seemed more important than interest rate policy in explaining the high inflation and unemployment episodes in recent US economic history. Regarding monetary policy, Primiceri (2005) found that the responses of the interest rate to inflation and unemployment exhibited a trend towards more aggressive behaviour, but this has had a negligible effect on the rest of the economy.

For some older EU countries, Ciccarelli and Rebucci (2006) used a TVC technique to study changes in the monetary transmission mechanism in the period 1981-1998. They found that the long-run impact on output

[1] For example, the vector autoregression (VAR) methodology is perhaps the most common econometric tool for the study of the transmission mechanism to key macroeconomic variables (see Christiano, Eichenbaum and Evans, 1999), which is a linear model.

of a common monetary policy shock decreased after 1991 in all countries studied. Using a more recent sample period, 1980-2007, Boivin, Giannoni and Mojon (2008) found that the creation of the euro has contributed to an overall reduction in the effects of monetary shocks, but also to a greater homogeneity of the transmission mechanism across countries. Using a structural open-economy model they argue that these findings can be attributed to the combination of a change in the policy reaction function (mainly towards more aggressive response to inflation and output) and the elimination of an exchange rate risk.

However, Weber, Gerke and Worms (2011) conclude that the empirical evidence on whether euro-area monetary transmission has changed is mixed. By adopting a fixed-parameter vector autoregression and endogenously searching for break dates, they conclude that euro-area monetary transmission after 1998 is not very different from before 1996, but probably very different compared to the interim period. They argue that this interim period could be responsible for the conflicting findings on whether monetary transmission has changed. In a more recent paper Cecioni and Neri (2011) find that euro-area monetary transmission has not significantly changed in time according to the estimation of a Bayesian vector autoregression before and after 1999. However, the estimation of their dynamic stochastic general equilibrium (DSGE) model over two subsamples indicated that euro-area monetary policy became more effective as a result of a decrease in the degree of nominal rigidities and a shift in monetary policy towards inflation stabilisation.

In this paper we study the transmission mechanism of monetary shocks to macroeconomic variables in three central European countries, the Czech Republic, Hungary and Poland, using quarterly data for 1993-2011. We also study the transmission mechanism of the euro area for comparison. The method we use is time-varying coefficient vector autoregression (TVC-VAR) identified through contemporaneous restrictions. We find that the transmission mechanism has changed in both the euro area and the three central European countries, and discuss various factors that may have contributed to these changes.

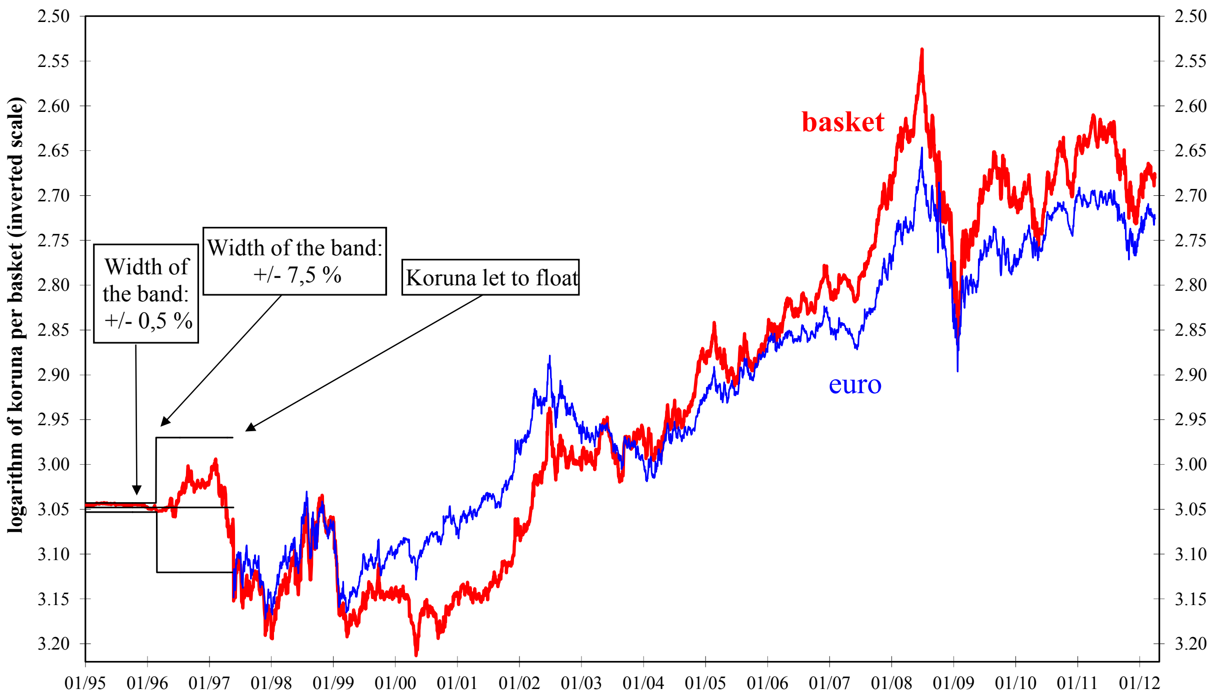
The rest of the paper is organised as follows. Section 2 briefly describes monetary regimes in the three central European countries. Section 3 details the methodology we use, which is followed by the description of the data in Section 4. The results of our econometric analysis are presented in section 5. Section 6 discusses the results. Finally, Section 7 presents a brief summary.

2 Monetary Regimes

The three central European countries had various kinds of exchange-rate targeting regime in the mid-1990s, which were changed gradually or suddenly to inflation targeting systems. The nominal exchange rate developments, which are shown in Figure 1, portray well these changes[2].

The Czech Republic had a narrow exchange rate band regime before 1996, when the band was widened to ±7.5%. Not much later, in May 1997, the band was swept away by a speculative attack, and the koruna was floated with occasional central bank intervention. Inflation targeting was introduced in January 1998. There has been just an implementation change since then: up-to 2001 year-end targets were defined and the fulfilment was analysed annually, while from 2002 continuous targets were pursued and the fulfilment was analysed monthly.

Figure 1.a: Nominal exchange rate of the Czech koruna, 1 January 1995 to 30 April 2012

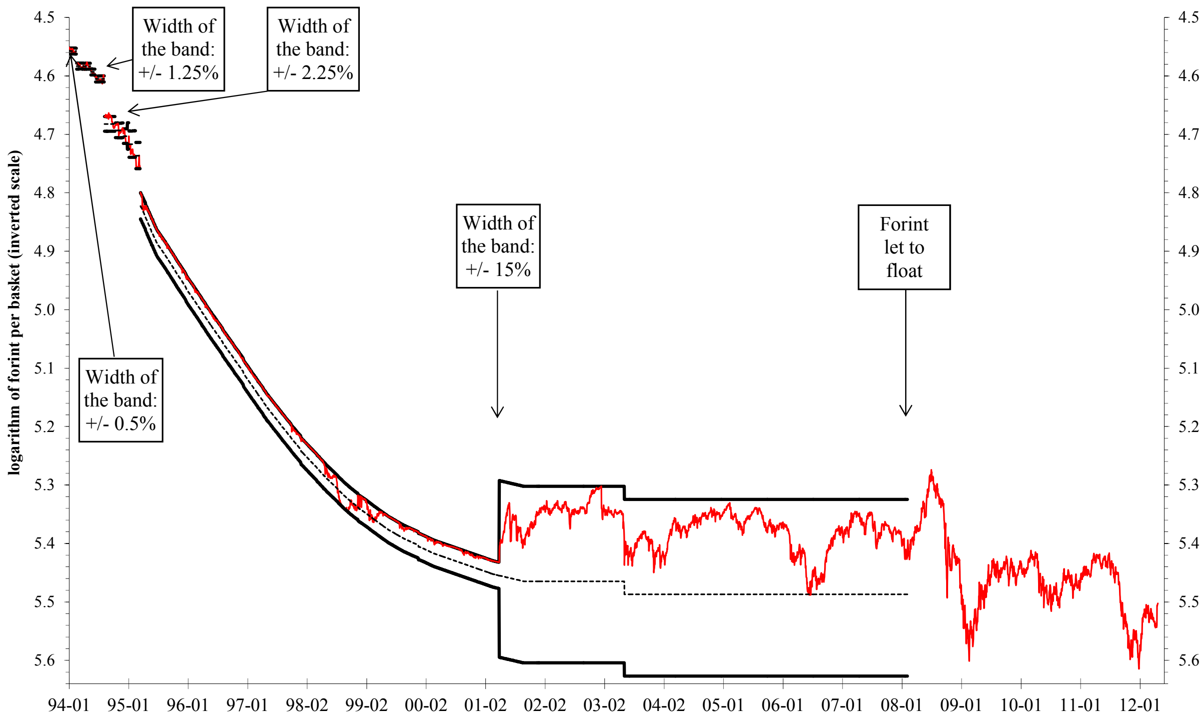


Source: Updated, using ECB data, from Darvas and Szapáry (2000).
Notes: Basket used prior to May 1997: 65% DEM and 35% USD. For better comparison, we show the rates against both the last basket and the euro for the floating period.

[2] For more information on the exchange rate systems of these countries, see Darvas and Szapáry (2010).

In Hungary, a pre-announced crawling band regime was introduced in March 1995 after a long period with an adjustable peg. The adopted band was narrow at ±2.25%. However, as Panel B of Figure 1 shows, the market rate eventually evolved like a crawling peg system, since it was almost continuously at the strong edge of the band (with the exception of the period of the Russian and Brazilian crises in late 1998 and early 1999). The band was widened substantially in May 2001 to ±15% and inflation targeting was introduced. However, at this time the exchange rate band and its crawling devaluation were kept. The rate of crawl was reduced to zero in October 2001. Following a strong appreciation pressure in early 2003, in June 2003 the band was unexpectedly devalued by 2%. In February 2008 the band was abandoned and free floating was introduced.

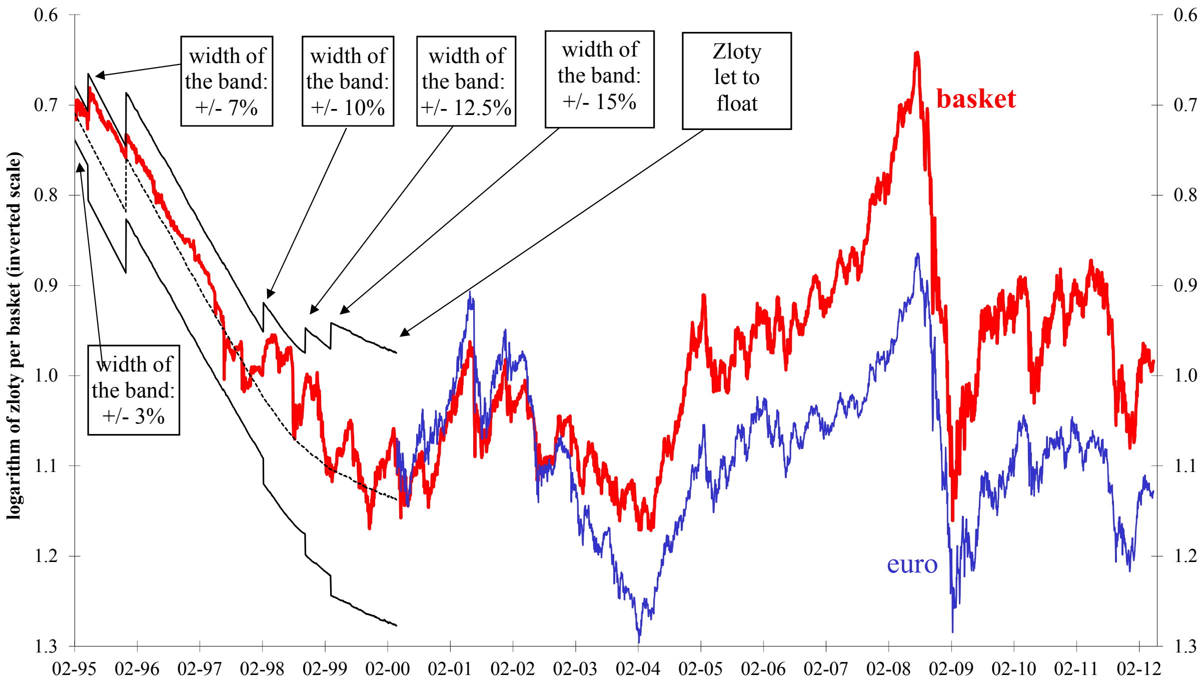
Figure 1.b: Nominal exchange rate of the Hungarian forint, 1 January 1994 to 30 April 2012



Source: Updated, using ECB data, from Darvas and Szapáry (2000).
Note. Composition of the basket: August 1993 - May 1994: 50% DEM and 50% USD; May 1994 - December 1996: 70% ECU and 30% USD; January 1997 - December 1998: 70% DEM and 30% USD; January - December 1999: 70% EUR and 30% USD; January 2000 - February 2008: 100% EUR. Forint was let to float in February 2008.

The Polish authorities made their exchange rate regime flexible gradually. As Panel C of Figure 1 indicates, Poland also adopted a pre-announced crawling band for most of the 1990s, but the band was widened to ±15% in several steps. There were heavy central bank interventions until 1997, which is also reflected in the relatively stable rate within the band, but since early 1998, the rate was allowed to move freely within the wide band. In April 2000, the band was abolished and inflation targeting was introduced with a freely floating exchange rate. Yet there were occasional central bank interventions, such as in late 2009, when many other floating exchange-rate countries also aimed to resist currency appreciation, or weaken their exchange rates[3].

Figure 1.c: Nominal exchange rate of the Polish zloty, 1 February 1995 to 30 April 2012



Source: Updated, using ECB data, from Darvas and Szapáry (2000).
Note: Composition of the basket prior to 1999: 45% USD, 35% DM, 10% GBP, 5% FRF, and 5% CHF; between January 1999 and April 2000: 55% EUR and 45% USD. Zloty was let to float in April 2000. For better comparison, we show the rates against both the last basket and the euro for the floating period.

These changes in monetary regimes most likely had an impact on the monetary transmission mechanism. This motivates our selection of time-varying parameter analysis, in addition to other structural changes and the potential inappropriateness of the linear approximation of the estimated model, as discussed in the introduction.

[3] See Darvas and Pisani-Ferry (2010).

3 Methodology

The econometric model we adopt is a time-varying coefficient vector autoregression with one exogenous variable. The reduced form of the model, which is estimated, can be written in the form:

$$(1) \qquad \mathbf{y}_t = \boldsymbol{\alpha}_t + \boldsymbol{\Phi}_t(L)\mathbf{y}_{t-1} + \boldsymbol{\gamma}_t x_t + \mathbf{v}_t \ ,$$

where $\mathbf{y}_t$ is the vector of endogenous variables, $\boldsymbol{\alpha}_t$ is the time-varying vector of intercepts, $\boldsymbol{\Phi}_t(L)$ is the time-varying parameter matrix polynomial in the lag operator *L*, $\boldsymbol{\gamma}_t$ is the time-varying parameter vector of the exogenous variable x_t, and $\mathbf{v}_t$ is the vector of estimated innovations. The time-varying nature of the parameter matrices and vectors makes the calculations more complicated. There are a number of issues to be addressed: the specification of time-varying parameters, lag length selection and specification, shock identification, impulse response analysis and initial conditions. The first four issues are discussed below, while the last is presented in the appendix.

3.1 Specification

Three main types of TVC-VAR models have been proposed recent years:

1) Parameters, treated as latent variables, assumed to follow random walks without drift and the Kalman filter is used for estimation[4]. This technique was used eg by Cogley and Sargent (2005), Primiceri (2005), Canova and Gambetti (2006), and Ciccarelli and Rebucci (2006)[5].

2) Parameters switch between regimes (back and forth) driven by a latent-state variable which follows a Markov switching process. This technique was used in eg Sims and Zha (2006).

3) Parameters change from one regime to another smoothly (and permanently) in time; the specification is the multivariate extension of the STAR (smooth transition threshold autoregression) model. This technique was developed in He, Teräsvirta and González (2009).

In this paper we follow the papers indicated in the first group by assuming that reduced-form VAR parameters follow driftless random walks and use the Kalman filter for maximum likelihood estimation and inference. The

[4] See Chapter 13 of Hamilton (1994) for an excellent exposition of Kalman-filtering and the maximum likelihood estimation of the parameters of a state-space system.

[5] To be more precise, Ciccarelli and Rebucci (2006) did not use a VAR but adopted a two-stage approach. In the first stage they measured monetary policy by estimating a system of reaction functions, and in the second stage they estimated output equations using the first stage estimates of monetary policy. Consequently, they do not model exchange rates and inflation rates.

main reason for our choice is that the random walk specification is a flexible model which can capture various time paths of the parameters. In contrast, the STAR-type time specification assumes a particular path and a smooth transition between the beginning and the end regime, which could be too restrictive for out purposes. The Markov switching specification, on the other hand, assumes a certain number of regimes having different (but fixed in time) parameters and that the process switches back and forth between these regimes. Specifying a limited number of states seems less attractive compared to the flexibility of the random walk specification in central European economies, which presumably did not oscillate between certain regimes, but have evolved gradually and may have not returned to an earlier regime. Nonetheless, it would be interesting to apply the other two techniques as well in a future work for the study of the transmission mechanism of the new member states.

3.2 Lag length

A key issue in TVC-VAR analysis is the selection of the lag length. Larger lags require the estimation of a large number of parameters, which is more difficult in a TVC than in a fixed parameter setting. When the order of the VAR is just one, for a four variable VAR, sixteen TVCs need to be estimated when the VAR does not include an intercept. This implies sixteen parameters; they are the standard errors of the innovations driving the parameters in the state equations. There are four additional parameters in the measurement equations as well; they are the standard errors of the innovations of the measurement equations. Hence, in a VAR(1) without an intercept 20 parameters need to be estimated. We found that a VAR(1) is not satisfactory with the exception of the Czech Republic: innovations in some equations were autocorrelated. Increasing the order of VAR without any constraint would lead to an explosion in the number of parameters to be estimated. Consequently, we adopt a certain type of restriction. Stock and Watson (2005), who estimate fixed parameter factor-structural vector autoregression (FSVAR) for the G-7 countries, allow more lags only for the left-hand side variable and restrict the number of lags of all other variables to one. We adopt this restriction to our TVC-VAR, which leads to the following specification:

$$
\begin{aligned}
\begin{bmatrix} y_{1,t} \\ y_{2,t} \\ y_{3,t} \\ y_{4,t} \end{bmatrix} &= \begin{bmatrix} \alpha_{1,t} \\ \alpha_{2,t} \\ \alpha_{3,t} \\ \alpha_{4,t} \end{bmatrix} + \begin{bmatrix} \phi_{11,t}^{(1)} & \phi_{12,t}^{(1)} & \phi_{13,t}^{(1)} & \phi_{14,t}^{(1)} \\ \phi_{21,t}^{(1)} & \phi_{22,t}^{(1)} & \phi_{23,t}^{(1)} & \phi_{24,t}^{(1)} \\ \phi_{31,t}^{(1)} & \phi_{32,t}^{(1)} & \phi_{33,t}^{(1)} & \phi_{34,t}^{(1)} \\ \phi_{41,t}^{(1)} & \phi_{42,t}^{(1)} & \phi_{43,t}^{(1)} & \phi_{44,t}^{(1)} \end{bmatrix} \begin{bmatrix} y_{1,t-1} \\ y_{2,t-1} \\ y_{3,t-1} \\ y_{4,t-1} \end{bmatrix} + \begin{bmatrix} \phi_{11,t}^{(2)} & 0 & 0 & 0 \\ 0 & \phi_{22,t}^{(2)} & 0 & 0 \\ 0 & 0 & \phi_{33,t}^{(2)} & 0 \\ 0 & 0 & 0 & \phi_{44,t}^{(2)} \end{bmatrix} \begin{bmatrix} y_{1,t-2} \\ y_{2,t-2} \\ y_{3,t-2} \\ y_{4,t-2} \end{bmatrix} + \cdots \\
&\cdots + \begin{bmatrix} \phi_{11,t}^{(p)} & 0 & 0 & 0 \\ 0 & \phi_{22,t}^{(p)} & 0 & 0 \\ 0 & 0 & \phi_{33,t}^{(p)} & 0 \\ 0 & 0 & 0 & \phi_{44,t}^{(p)} \end{bmatrix} \begin{bmatrix} y_{1,t-p} \\ y_{2,t-p} \\ y_{3,t-p} \\ y_{4,t-p} \end{bmatrix} + \begin{bmatrix} \gamma_{1,t} \\ \gamma_{2,t} \\ \gamma_{3,t} \\ \gamma_{4,t} \end{bmatrix} x_t + \begin{bmatrix} v_{1,t} \\ v_{2,t} \\ v_{3,t} \\ v_{4,t} \end{bmatrix}
\end{aligned}
\quad , \qquad (2)
$$

where $y_{i,t}$ denote the variables included in the VAR, $\alpha_{j,t}$, $\phi_{jk,t}^{(m)}$ and $\gamma_{j,t}$ denote the time-varying parameters to be estimated, and $v_{i,t}$ denote the innovations. We select the lag length on the basis of the Ljung-Box statistics to get rid of autocorrelation on the innovations, by allowing at most six lags as specified in equation (2).

3.3 Identification

A next issue is identification. Traditional identification, which usually takes the form of some contemporaneous and/or long-run restrictions, has been severely criticised. Contemporaneous restrictions have been criticised on the basis of being ad hoc, while long run restrictions have been shown to entail substantial estimation errors (eg in Faust and Leeper, 1997). A possible new way of identifying shocks is the recent sign restriction identification method (eg Uhlig, 2005; Canova and De Nicoló, 2002; and Peersman, 2005). Fry and Pagan (2011), on the other hand, highlight some drawbacks of the sign restriction methodology and argue for the superiority of contemporaneous restrictions. Due to the inconclusiveness of the debate over various identification techniques, in this paper we use contemporaneous restrictions to identify structural shocks, using the standard recursive information ordering (see section 11.4 of Hamilton, 1994). We order the endogenous variables as output, price, interest rate and exchange rate and thereby assume that monetary shock do not affect output and prices contemporaneously (we use quarterly data), but instantly affect the interest rate and the exchange rate. For better comparability of the impacts of monetary shock across countries, we normalise the impact to a 1 percentage point shock.

3.4 Impulse response analysis

The fourth issue is impulse response analysis. Since the model is nonlinear due to changing parameters, no unique impulse response function is available, but we can attach an impulse response function to each observation of the sample. Still, there are two possible ways to proceed. One could use either the parameter set of time *t*, or *t*, *t*+1, *t*+2, Ɛ (that is, take into account future parameter changes), to calculate the impulse response function.

Since parameters are assumed to follow driftless random walks, their forecasts are equal to their current values. Therefore, the first option has the rationale that is it based on the parameters 'known' at time *t* (to be more precise: estimated for time *t*), which are expected to apply to the future as well. Obviously, at the last observation of the sample only the first option can be used, at the last but one observation the second option can be used only for one observation, and so on. We used the first option in the whole sample period. In practice, we calculated the impulse responses as the difference between two simulations, a baseline one and one with a shock at time *t*, which is the standard way to derive impulse responses.

Impulse response functions are usually graphed to show the effects of a one standard deviation shock. Since the volatility of monetary shocks vary across countries and we also aim to compare the results across countries, we normalise the impulse response functions to show the effects of a 1 percentage point monetary policy shock.

4 Data

Our data set includes quarterly observations in the period from 1993Q1 to 2011Q4 for the four standard endogenous variables that are typically used for monetary transmission VARs of open economies: output, price, interest rate and exchange rate.

For the euro area we use the aggregate of the first 12 countries that introduced the euro in 1999/2001 in order to have a clear distinction between these 12 older EU member states and the member states that joined the EU in 2004 and later the euro area. Data is available for the aggregate of these 12 euro-area countries for the output, price and exchange rate measures we use, but not for the 3-month interest rates, which is for the whole euro area.

We use seasonally adjusted constant price GDP as a measure of output. The main data source is Eurostat which publishes data since 1995 (1996 for the Czech Republic). For the three central European countries, earlier data starting in 1993 was proxied using the method outlined by Várpalotai (2003), which makes use of the information in the available annual GDP figures and quarterly indicators, such as industrial output and retail sales, which were chained backward to Eurostat data. Euro-area GDP data, which is also available from Eurostat from 1995, was chained to data from the OECD that is available for earlier years as well.

For measuring prices we did not use the all-items consumer price index, because it includes many volatile items that are not much affected by monetary policy, but serve only to introduce noise into consumer price statistics. Instead, we used the Eurostat measure 'Overall index excluding energy, food, alcohol and tobacco' (OOXEFOOD) from the harmonised index of consumer prices (HICP) database, which is available for our full sample period for the euro area, but only since 1996 for the central European countries. For the earlier years we used the core inflation measure calculated by Darvas (2001) that is based on aggregating detailed consumer price statistics from national statistical offices. In the overlapping period for which both OOXEFOOD and the core inflation measure of Darvas (2001) are available, the two series are highly similar for all countries studied. We seasonally adjusted the resulting price level series with X12.

The interest rate we use is the three month interbank interest rate taken from Eurostat. Data for Hungary is available from 1994 and for Poland from 1995, therefore, we augmented these series using data from the central banks for earlier years.

As the measure of exchange rate, we use the nominal effective exchange rate calculated against the basket of 138 trading partners from Darvas (2012).

The exogenous variable we use is the commodity price index published in the IMF's International Financial Statistics.

Output, price, exchange rate and commodity price index are included as logarithmic first differences. The interest rates are included in levels, or to be more precise, as $\ln(1+i_t)$, in order to be consistent with the logarithmic first differences of the other variables. We have also removed the mean of these series before estimating the TVC-VAR models[6].

5 Results

We allow at most six lags in the VAR (using the approach described in Section 3.2) to limit the number of parameters to be estimated. For the same reason, and because our data is demeaned before estimation, we do not wish allow a time-varying intercept. However, in the case of Hungary, the model without the intercept turned out to produce unrealistic results. We therefore added a time-varying intercept for Hungary, which is motivated by the fact that inflation was rather persistent and declined only gradually, and even picked-up significantly a number of times during our sample period. The time-varying intercept can capture such development.

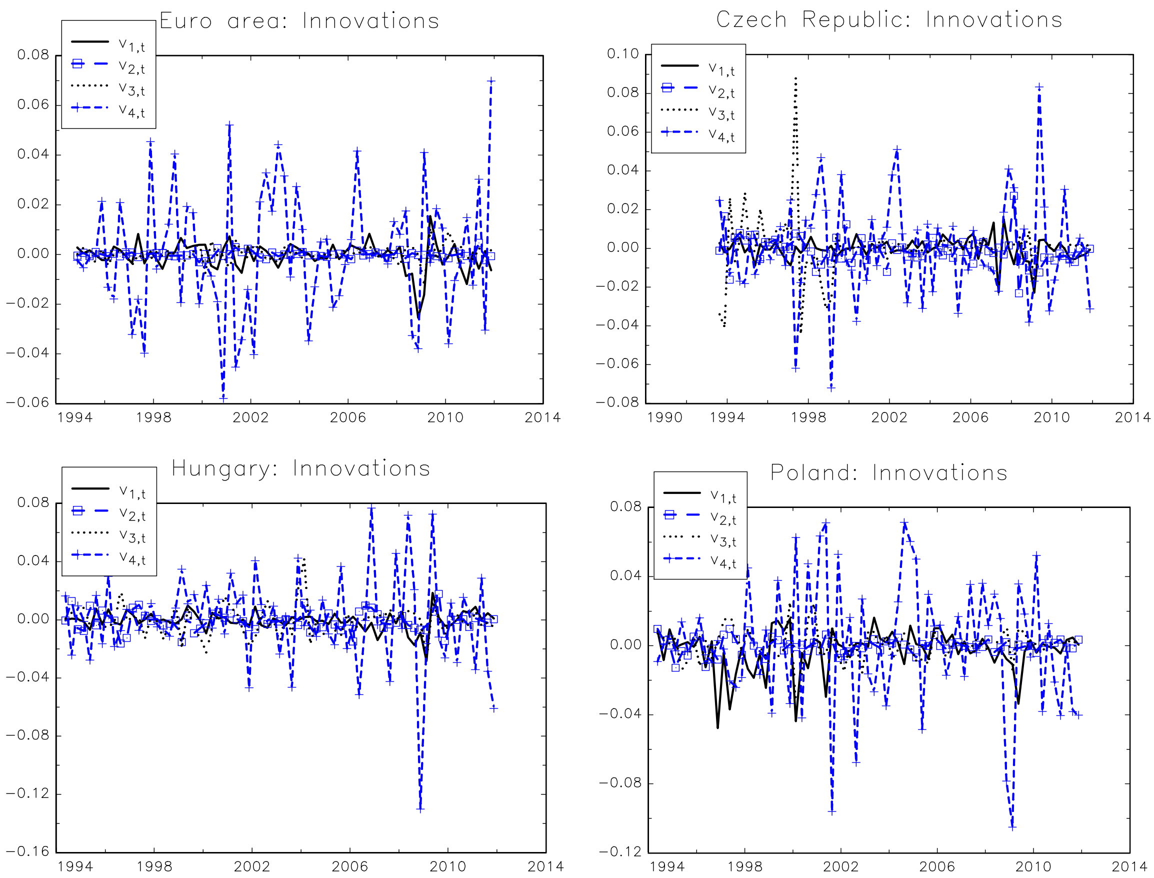
We selected the lag length to eliminate, to the largest possible extent, autocorrelation of the innovations of the equations. Six lags for the euro area, four lags for Hungary and Poland and one lag for the Czech Republic were the best choices. Figure 2 shows the innovations. Table 1 indicates that for the Czech Republic, the null hypothesis of no autocorrelation can be rejected for the first innovation, but as we extended the lag length to higher levels, test statistics became worse and therefore we chose only one lag.

Table 2 shows that the null hypothesis of normality can be rejected for most innovations, and hence the

[6] To completely get rid of non-zero means, we removed the mean separately from the left and right hand side variables. For example, when the lag length is two, the first variable and its lags were mean-removed as: $\tilde{y}_{1,t} = y_{1,t} - \sum_{\tau=3}^{T} y_{1,\tau} \Big/ (T-2)$, $\tilde{y}_{1,t-1} = y_{1,t-1} - \sum_{\tau=2}^{T-1} y_{1,\tau} \Big/ (T-2)$, and $\tilde{y}_{1,t-2} = y_{1,t-2} - \sum_{\tau=1}^{T-2} y_{1,\tau} \Big/ (T-2)$, where *T* denotes the number of observations before taking lags.

estimate can be regarded as quasi maximum likelihood estimates. Note that, as Hamilton (1994, section 5.8) highlights, maximum likelihood estimation may still be a reasonable way to estimate parameters even if the data were not generated by the assumed density.

Figure 2: Innovations



Note. The four series shown are innovations to the equations of output growth, inflation, interest rate and nominal effective exchange rate changes, respectively.

Table 1: Ljung-Box autocorrelation test for innovations

	Euro area	Czech Rep.	Hungary	Poland
$v_{1,t}$	7.2	17.8	7.9	7.8
	(0.408)	(0.013)	(0.341)	(0.351)
$v_{2,t}$	6.7	5.5	7.2	7.3
	(0.459)	(0.599)	(0.410)	(0.403)
$v_{3,t}$	11.8	6.6	9.0	13.0
	(0.106)	(0.467)	(0.249)	(0.071)
$v_{4,t}$	6.3	10.4	10.8	7.8
	(0.501)	(0.166)	(0.149)	(0.355)

Note. P-values in parentheses.

Table 2: Jarque-Bera normality test for innovations

	Euro area	Czech Rep.	Hungary	Poland
$v_{1,t}$	79.4	94.8	64.3	50.9
	(0.000)	(0.000)	(0.000)	(0.000)
$v_{2,t}$	21.1	36.1	0.4	2.4
	(0.000)	(0.000)	(0.821)	(0.295)
$v_{3,t}$	18.3	544.5	29.1	19.6
	(0.000)	(0.000)	(0.000)	(0.063)
$v_{4,t}$	0.6	8.1	37.7	2.3
	(0.753)	(0.017)	(0.000)	(0.322)

Note. P-values in parentheses.

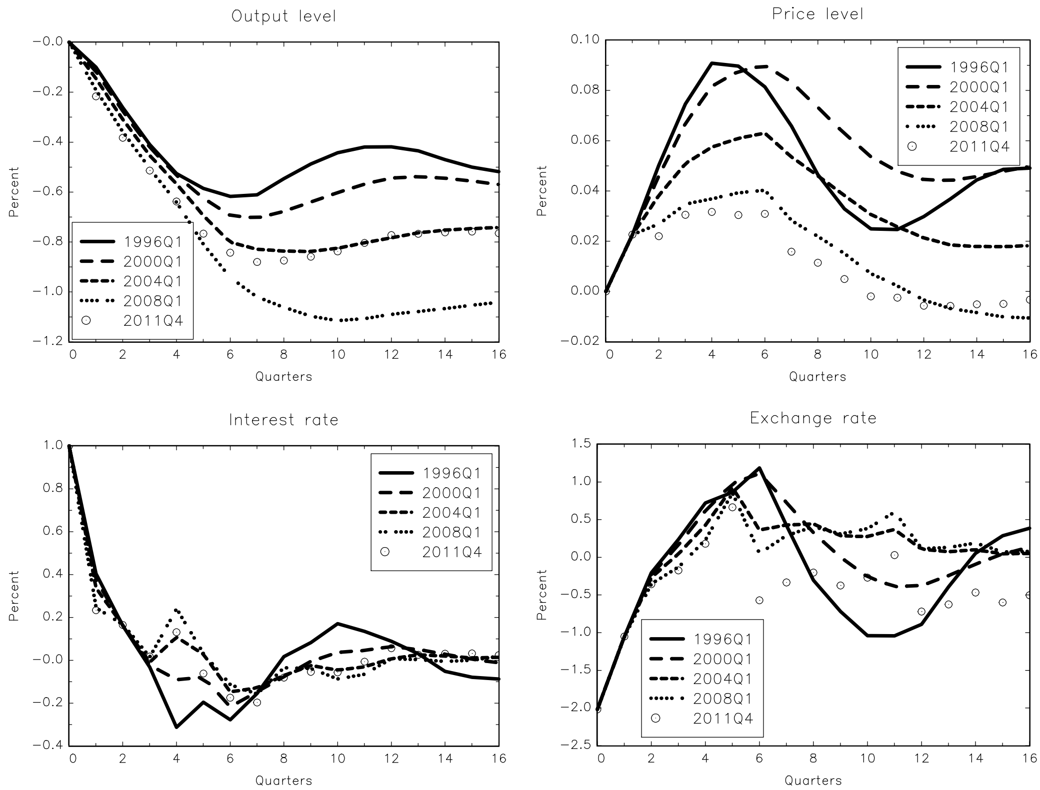
Due to the time varying parameters, impulse response functions also change from quarter to quarter. In order to conserve space and to enhance the discussion of our results, in Figures 3, 4, 5 and 6 we show impulse response functions for five different dates of our sample: 1996Q1, 2000Q1, 2004Q1, 2008Q1 and 2011Q4[7]. These dates are equidistant (the distance between the last two ones is one quarter less than the rest, due to data availability) and can well portray the changing nature, if any, of impulse responses. We did not select the exact dates when the monetary regimes changed, for four reasons. First, these dates differ across countries, which would make it more difficult to compare the results. Second, Hungary and Poland had multiple changes,

[7] The impulse responses were calculated on the basis of the time-varying parameters derived from the Kalman-smoother.

while the Czech Republic has not had a major change since 1998. Third, monetary transmission may not change instantly at the time of regime change. And forth, we are interested in studying changes in monetary transmission that happen for reasons other than the monetary regime change.

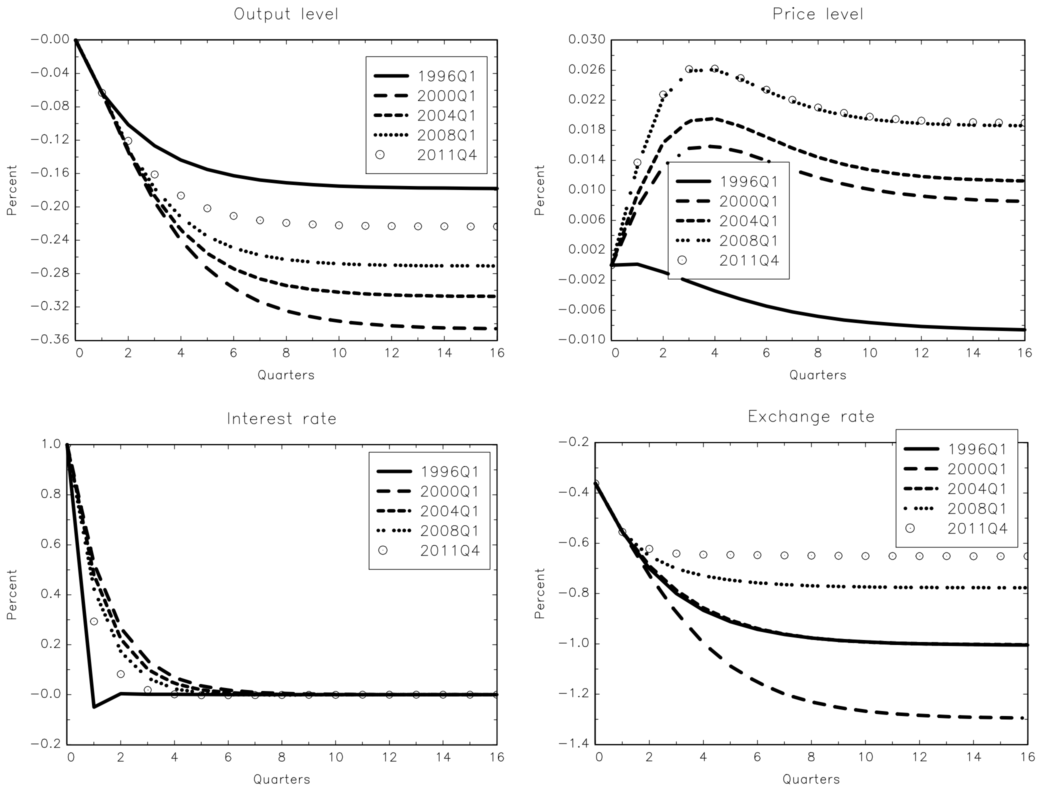
Figures 3, 4, 5 and 6 indicate that impulse responses to a monetary shock have changed in time in the euro area and also in the three central European economies. Yet the changes are not monotonic.

Figure 3: Euro area: Impulse response functions at different dates



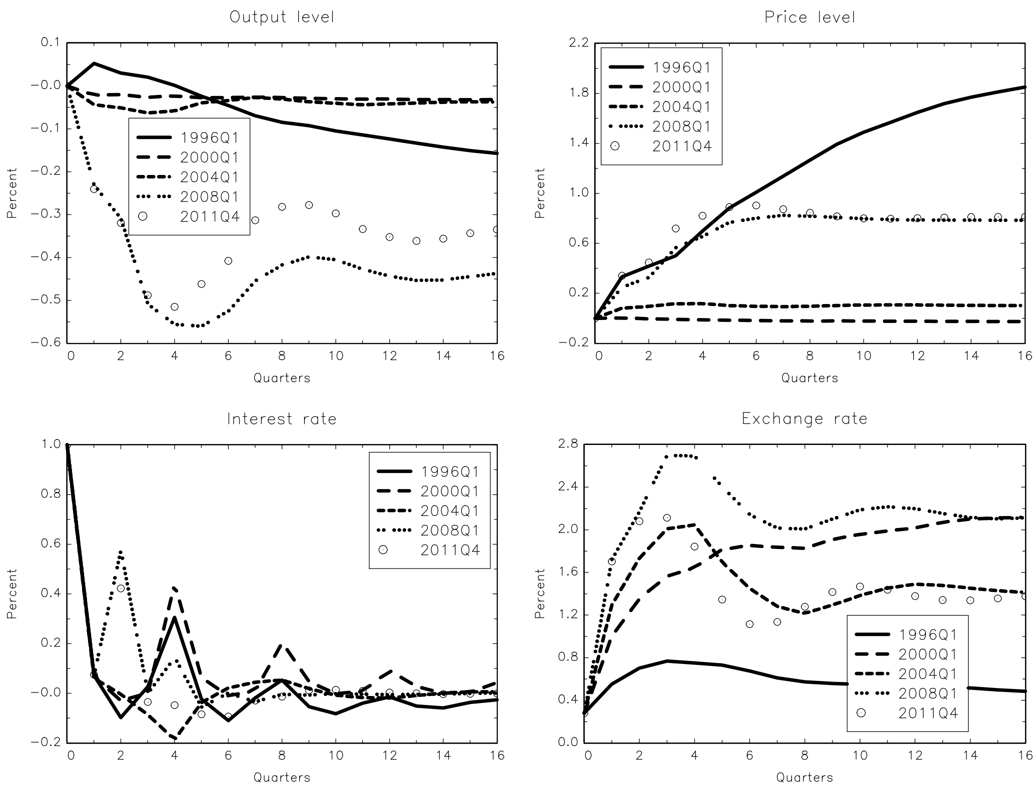
Note: the size of the shock is 1 percentage point.

Figure 4: Czech Republic: Impulse response functions at different dates



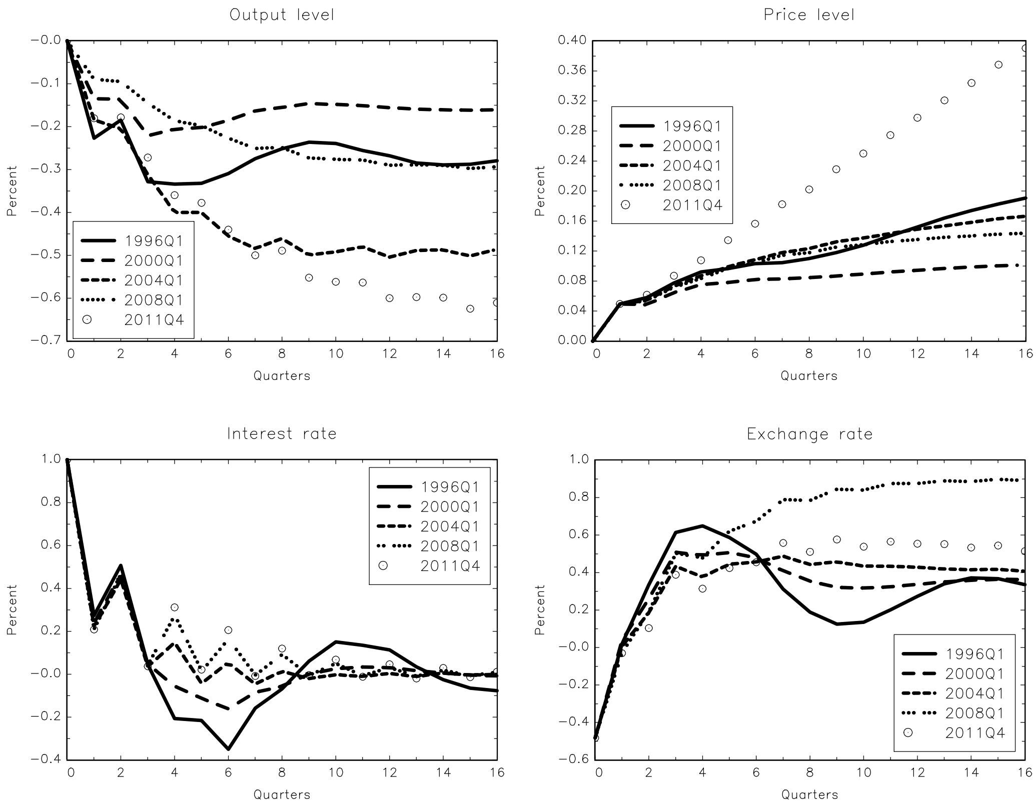
Note: the size of the shock is 1 percentage point.

Figure 5: Hungary: Impulse response functions at different dates



Note: the size of the shock is 1 percentage point.

Figure 6: Poland: Impulse response functions at different dates



Note: the size of the shock is 1 percentage point.

In the euro area, the response of output was broadly the same in 1996 and 2000: a one percentage point contractionary monetary shock lowered the long-run level of output by about 0.6 percent. Then impact gradually increased until 2008 to about 1.0 percent. However, by 2011 the impact is estimated to be lower, 0.8 percent, which is broadly equal to the 2004 impact, suggesting that monetary policy became less effective during the global financial and economic crisis. For prices we got the so called 'price puzzle' result, that is, prices increase after a monetary contraction, even though we have included commodity prices as an exogenous variable. Note, however, that the magnitude of price response is extremely low: a one percent contractionary monetary shock is estimated to increase the price level by a maximum of 0.09 percent (in 1996 and 2000) after about a year, and even less at longer horizons and later years. For example, in 2008 and 2011 this tiny price puzzle diminishes as the long run impact on prices turns to negative.

The price puzzle result is a general finding of many monetary VARs. The reason is that monetary policy used to tighten when prices are expected to increase. But if the monetary tightening is not able to fully eliminate the rising inflation, then estimated results will indicate that prices increase after a monetary tightening, even if the tightening lowered the actual price increase. The standard approach to get rid of this result, following a suggestion by Sims (1992), is to include commodity prices in the VAR as an indicator of future inflation. Hanson (2004) argues that the ability of commodity prices to mitigate the price puzzle may be due to an information channel: commodity prices respond more quickly than aggregate goods prices to future inflationary pressures. The inclusion of the index of commodity prices in our specification reduced the price puzzle result for the euro area, but did not eliminate it completely.

The responses of the euro-area interest rate and exchange rate did not change much in time, suggesting that structural changes occurred mostly in the real economy.

Figure 4 shows that the changes of the shapes of the output response were also not monotonous in the Czech Republic. The impact of a one percentage point monetary shock was the lowest, 0.16 percent, in 1996 when a fixed exchange rate regime was in place. Then the response increased to 0.35 by 2000 and declined afterward, including during the global financial and economic crises, to 0.22 in 2011. Our results are rather similar, both in terms of time-variation and magnitude, to the findings of Franta, Horváth and Rusnák (2011), even though they use a different methodology. According to Figure 2 of Franta, Horváth and Rusnák (2011), in 1996 the output response to a one percentage point contractionary monetary shock was about 0.18 percent after three years, which more or less gradually increased to about 0.35 in 2006, and then reduced to about 0.25 by 2010. They estimate the model with Bayesian techniques, allow time-variation of the variances of the shocks and identify the shocks with a mixture of contemporaneous and sign restrictions, and consequently argue that their methodology is superior to my methodology. Yet the similarity of the results suggests that these methodological differences are not that important.

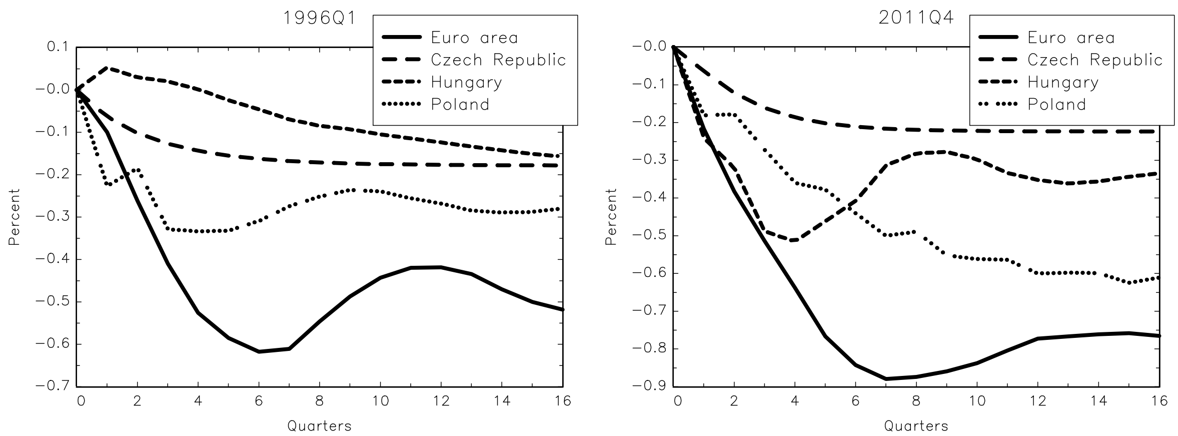
We also find a minuscule price puzzle for the Czech Republic, ie a one percent monetary shock increases the price level by 0.02 percent. This finding is different from the results of Franta, Horváth and Rusnák (2011) and may be related to the different shock identification method they used.

In Hungary monetary disturbances had only a tiny effect on output in the period 1996-2004, when the exchange rate was tightly managed, but the effects increased by 2008 to about 0.44 percent, when the freely floating exchange rate regime was introduced (Figure 5). Similarly to the Czech Republic and the euro area, the effect of a monetary shock on Hungarian output declined somewhat between 2008 and 2011. At the last observation of our sample the long-run impact of a one percentage point monetary shock in output is estimated to be 0.34 percent. The price-puzzle is a strong feature of the results.

Monetary policy generally became more effective in Poland, as it is shown by Figure 6, but the variation across time was not monotonous here as well. The long run effect of a 1 percentage point monetary shock increased from around 0.15-0.30 percent in 1996 and 2000 to around 0.5 percent by 2004. However, by 2008 the output impact is estimated to decline, but increased again by 2011 to about 0.6 percent. The price puzzle also characterises our results for Poland, though to a lesser extent than the results for Hungary.

Figure 7 compares output responses in each of the three central European countries with that of the euro area in 1996Q1 and 2011Q4, normalised, again, to a 1 percentage point monetary shock. In 1996, monetary policy was the least effective in Hungary, though the impact in the Czech Republic was just a little greater. The impact in Poland was also limited, though. By 2011 the impact increased in all three countries and the euro area, and in the ordering only the Czech Republic and Hungary changed places. The strength of the Polish response is just slightly lower than in the euro area, though the time profile of the response is somewhat different, since in Poland the impact takes place more gradually.

Figure 7: Output response to a monetary shock in 1996Q1 and 2011Q4



Note: the size of the shock is 1 percentage point.

6 Discussion

Differences in monetary transmission across countries could be explained by several factors. Unfortunately, due to the small number of countries we study we cannot perform a formal econometric analysis to link the strength of monetary transmission to its possible determinants. Therefore, in this section, we can only assess certain factors suggested in the literature and other factors that we regard as important.

Georgiadis (2012) formally analysed the differences in monetary transmission focusing on several OECD countries. He concluded that the financial structure, labour market rigidities and the industry mix account for a large share of cross-country differences in the output response to a monetary shock. However, as we show below, these factors do not associate well with the strength of monetary transmission in the countries we studied.

Concerning financial structure, Georgiadis (2012) considers the degree of competition in the banking sector (higher competition leads to faster and more complete pass-through of policy interest rates to retail interest rates, thereby making monetary policy more powerful) and the importance of bank credit in the economy (higher importance increases the role of the credit channel[8]). To measure competition he uses the net interest margin of banks. Table 3 indicates that the association of net interest margin with the strength of monetary transmission is far from perfect in the sample of our countries. While the relationship between the euro area, Poland and Hungary are in line with expectations, this is only after 2001, and the results for the Czech Republic are very much against the supposed relationship. The Czech Republic had the lowest net interest margins among the three central European countries, and her values were similar to, or even lower than, the euro area's. Yet according to our results, the strength of the monetary transmission is much weaker in the Czech Republic than in the euro area and in Poland, and also than in Hungary after 2008. Also, the net interest margin increased in Hungary recently, yet monetary transmission is also estimated to be more powerful. It is interesting to observe that the US, which is also included in Table 3 for comparison, has values well above the values of the euro area, the United Kingdom and the Czech Republic.

Table 3: Net interest margin (multiplied by 100), 1995-2009

	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009
Czech Republic	3.2	2.3	3.0	4.8	2.8	2.2	2.1	2.0	2.0	2.1	2.7	2.2	2.5	2.8	3.2
Hungary	5.2	4.4	3.8	3.6	3.5	4.4	4.9	5.8	5.5	5.6	6.0	5.5	6.1	6.9	7.9
Poland	7.3	7.0	6.2	6.5	5.1	4.8	4.0	5.1	4.3	4.0	4.3	4.4	3.7	3.1	2.5
Euro area 12	3.1	3.0	2.9	3.0	2.8	3.3	2.9	2.6	2.5	2.6	2.3	2.4	2.2	2.2	2.1
United States	4.4	4.4	4.5	4.4	4.2	4.3	4.1	4.1	4.0	4.0	4.0	4.0	3.9	3.8	3.7
United Kingdom	2.2	2.3	2.8	2.9	2.8	3.2	2.3	2.0	2.4	2.7	2.9	2.9	2.0	1.4	0.9

Source: Beck and Demirgüç-Kunt (2009). The aggregate for euro area 12 was calculated by the author by weighting country-specific data with shares in euro-area GDP.
Note: net interest margin = accounting value of bank's net interest revenue as a share of its interest-bearing (total earning) assets.

Another aspect considered by Georgiadis (2012) is the importance of credit in the economy. While we found the strongest impact of monetary policy in Poland among the three central European countries, the share of credit to the private sector in GDP was lower than in the other two countries before 2008, and by 2008 it has just increased to the Czech values (Table 4).

Table 4: Credit to the private sector by resident banks (percent of GDP), 1995-2010

	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009	2010
Czech Republic	71	69	70	61	56	49	41	31	32	33	37	42	49	54	55	53
Hungary	22	21	24	23	25	32	33	35	43	46	51	57	64	73	71	70
Poland	17	19	21	23	26	27	27	27	28	28	29	35	41	52	53	55
Euro area 12	82	83	86	90	93	98	100	100	102	104	109	114	120	125	131	132
United States	47	48	48	49	49	50	52	52	53	55	57	59	61	61	64	65
United Kingdom	113	117	117	116	118	129	134	138	143	151	160	171	187	211	214	210

Source: Author's calculation using data from the IMF International Statistics. The aggregate for euro area 12 was calculated by the author by weighting country-specific data with shares in euro-area GDP.

[8] See Égert and MacDonald (2009) for a nice exposition of the credit channel.

Furthermore, the presence of foreign currency loans blurs the assessment of the credit stock. Central bank interest rate changes do not directly translate into interest rate changes of foreign currency loans, even if the loan interest rates are variable, since the main determinant of the foreign currency interest rate is the money market rate of the foreign currency. On the other hand, domestic interest rate changes could influence the exchange rate and cause balance sheet effects through the revaluation of foreign currency loans. We found for both Hungary (Figure 4) and Poland (Figure 5) that a monetary shock tends to appreciate the exchange rate of the local currency and thereby create a positive wealth effect (by reducing the local currency value of foreign currency loans). This positive wealth effect could even boost domestic demand through the credit channel, instead of contracting it, after an unexpected interest rate increase by the central bank[9].

As Table 5 indicates, there were marked differences between the three central European countries in terms of the share of foreign currency loans in total loans. The share has increased to 65 percent in Hungary and about 30 percent in Poland, while in the Czech Republic it has remained stable at about 15 percent. The high share in Hungary has significantly constrained the effectiveness of domestic monetary policy.

Table 5: Share of foreign currency loans in private sector credit (percent), 1996-2011

	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009	2010	2011
Czech Republic	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	16	14	13	13	14	13	14	13	14	15
Hungary	n.a.	n.a.	28	29	33	30	29	34	39	46	50	57	65	64	64	63
Poland	13	16	22	20	22	25	27	30	24	25	26	24	33	31	31	33

Source: author's calculation using data from Czech National Bank, Central Bank of Hungary and National bank of Poland.

Another aspect considered by Georgiadis (2012) is labour market rigidities, using the OECD's index of strictness of employment protection. Labour market rigidities may dampen the sensitivity of inflation to monetary shocks and enhance that of output. This indicator does not associate again perfectly with our findings for the strength of monetary transmission. While we found that Poland had the most powerful transmission among the three countries and, as Table 6 indicates, it had the most restrictive employment protection staring in 2003, in earlier years protection was less stringent. Also, the euro area has had a higher level of protection than any of the three central European countries, yet the strength of transmission was estimated to be higher in all years.

[9] Moreover, panel econometric analysis based on data for the Czech Republic, Hungary, Poland and Slovakia of Brzoza-Brzezina, Chmielewski and Niedźwiedzińska (2010) showed that restrictive monetary policy leads to a decrease in domestic currency lending, but simultaneously accelerates foreign currency denominated loans, which complicates the conduct of monetary policy.

Table 6: Strictness of employment protection index, 1995-2008

	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008
Czech Republic	1.94	1.94	1.94	1.94	1.94	1.94	1.94	1.94	1.94	1.94	2.10	2.10	1.99	1.99
Hungary	1.54	1.54	1.54	1.54	1.54	1.54	1.54	1.54	1.75	1.75	1.75	1.75	1.85	1.85
Poland	1.86	1.86	1.86	1.86	1.86	1.86	1.86	1.65	2.07	2.19	2.19	2.19	2.19	2.19
Euro area 11	3.21	3.21	2.87	2.76	2.73	2.71	2.65	2.59	2.54	2.54	2.54	2.54	2.54	2.53
United States	0.65	0.65	0.65	0.65	0.65	0.65	0.65	0.65	0.65	0.65	0.65	0.65	0.65	0.65
United Kingdom	0.98	0.98	0.98	0.98	0.98	1.05	1.05	1.10	1.10	1.10	1.10	1.10	1.10	1.10

Source: OECD Strictness of Employment Protection Legislation dataset. The aggregate for euro area 11 (euro area 12 excluding Luxembourg, due to missing data), was calculated by the author by weighting country-specific data with shares in euro-area GDP.
Note: The OECD indicators of employment protection measure the procedures and costs involved in dismissing individuals or groups of workers and the procedures involved in hiring workers on fixed-term or temporary work agency contracts. The index ranges from zero (least stringent protection) to six (most restrictive protection). For 1998-2008 we show version 2 of this indicator, which is based on 18 data items on regular employment, temporary employment and collective dismissals and is available for 1998-2008. For 1995-97 we have chained backward version 1 of this indicator, which is based on 14 data items on regular employment and temporary employment.

The third item considered by Georgiadis (2012), the share of durable goods in manufacturing output, also does not associate well with our findings. According to Table 4 of his paper, this share was about 15 percent in the Czech Republic, 13 percent in Hungary and about 9 percent in Poland. The average share for the euro area could be about 8-9 percent. Yet we found that in Poland and the euro area, monetary transmission was much more powerful than in the Czech Republic and Hungary.

Yet there are other factors that could explain the strength of monetary transmission. There is a very obvious factor, the exchange rate regime, which is clearly supported by our analysis. In exchange rate targeting regimes, monetary transmission was weaker than in inflation targeting regimes with floating exchange rates in all three central European countries we studied.

Another aspect could be related to the credibility of monetary policy. As discussed in Section 2, Hungarian monetary policy adopted a crawling peg regime from 1995 to 2001. The approach of the central bank was to tightly manage the exchange rate in order to gradually decrease inflation without risking large trade imbalances, which was, by the way, a rather successful strategy for achieving these goals[10]. While the widening of the band and the shift to inflation targeting in 2001 indicated a change in policy preferences, the tiny 2 percent devaluation compared to the ±15 percent width of the band in 2003 signified that monetary authorities had changed their preferences, largely weakening the credibility of the claim that pure inflation targeting is done. Up to 2003, the exchange rate had been allowed to reach the strong edge of the band fuelled by revaluation expectations. The devaluation in 2003 signalled that the authorities no longer wanted to allow a strong nominal exchange rate. In contrast, the abandonment of the exchange rate band in 2008 could have indicated that inflation will receive more attention in the future. Largely as a consequence of policy preferences, inflation was rather high in Hungary during the 2000s (Table 7). These factors could have

[10] See more details in Szapáry and Jakab (1998).

contributed to the low credibility of central bank policy, which might result in the low market response to monetary actions. The finding that response increased somewhat by 2008 and 2011 (Figure 5) could indicate that the credibility of monetary policy had increased by 2008, perhaps due to the abolishment of the exchange rate band and the introduction of the floating exchange rate.

Poland, on the other hand, adopted a very tight monetary policy even in an economic downturn in the early 2000s, when the real interest rate even exceeded 10 percent per year (Table 8). This could have contributed to the high credibility of monetary actions: market participants learned that policy was strict and resolute. Credibility of Czech monetary policy could have been in between of Hungary and Poland. The Czech central bank did not adopt such an aggressive monetary policy as Poland around the turn of the millennium but nonetheless it was very successful, since the average inflation was the lowest among the three central European countries in our sample period, and in some years inflation was even lower than in the euro area[11].

Table 7: Core inflation (annual percent change), 1995-2011

	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009	2010	2011
Czech Republic	8.2	6.7	6.4	6.8	1.6	1.5	3.2	2.5	0.6	1.9	1.0	0.9	1.6	4.0	0.4	0.0	-0.1
Hungary	24.7	23.1	15.8	13.2	9.9	7.6	8.7	5.6	4.8	5.7	3.3	1.9	5.4	3.2	4.0	3.0	1.7
Poland	29.2	21.3	15.7	13.7	10.0	9.2	5.7	2.7	1.0	1.9	1.0	0.5	1.3	2.1	2.4	1.3	2.0
Euro area 12	2.8	2.3	1.6	1.5	1.1	1.0	1.8	2.4	1.8	1.8	1.4	1.4	1.9	1.8	1.4	1.0	1.4
United Kingdom	n.a.	n.a.	1.9	1.4	0.7	0.1	1.1	1.5	1.2	1.0	1.4	1.3	1.6	1.6	1.7	2.7	3.0

Source: see section 2.

Table 8: Real three-month interest rate (percent), 1995-2010

	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009	2010
Czech Republic	4.0	5.3	8.6	12.5	5.3	2.1	2.6	2.9	0.4	1.4	1.1	0.7	-0.9	3.7	2.2	1.4
Hungary	6.7	7.3	6.4	7.3	7.0	2.5	5.0	4.2	2.6	7.8	5.1	1.5	4.4	4.5	6.1	4.4
Poland	5.2	4.9	8.8	9.5	5.0	12.3	13.0	7.9	3.7	5.2	4.8	2.9	2.5	3.8	3.1	1.8
Euro area 12	4.4	3.5	2.8	2.8	1.9	2.5	1.8	1.5	0.5	0.7	0.8	1.2	2.4	3.2	0.2	-0.6
United Kingdom	n.a.	4.2	5.4	6.7	5.5	5.1	3.5	2.8	2.7	3.2	3.4	3.2	4.4	3.8	-1.4	-2.2

Source: author's calculation using data detailed in section 2.
Note: the real interest rate was calculated using actual future core inflation, ie the 1995 nominal interest rate is deflated with actual core inflation in 1996.

Finally, openness could also explain the differences, since the Czech Republic and Hungary are much more open than Poland (Table 9). In more open economies output developments depend more on the foreign business cycle than in less open economies. This feature could mitigate the effects of domestic monetary policy.

[11] Jarocinski (2010), who uses a fixed-parameter panel VAR identified with a mixture of contemporaneous and sign restructons, also hypothesize that the level and variablility of inflation may have a role in determining the strenght of monetary transmission.

Table 9: Exports of goods and services (percent of GDP), 1995-2011

	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009	2010	2011
Czech Republic	48	47	50	52	53	61	63	58	59	63	64	67	68	64	60	68	74
Hungary	45	49	55	62	65	75	72	63	61	63	66	78	81	82	78	87	93
Poland	23	22	23	26	24	27	27	29	33	37	37	40	41	40	39	42	43
Euro area 12	29	30	32	33	33	37	37	36	35	36	38	40	41	42	36	40	43
United States	11	11	12	11	11	11	10	9	9	10	10	11	12	13	11	13	14
United Kingdom	28	29	29	27	26	28	27	26	25	25	26	29	27	29	28	30	32

Source: Ameco. The aggregate for euro area 12 was calculated by the author by weighting country-specific data with shares in euro-area GDP.
Note: The aggregate for the euro area also considers intra-euro area trade, while, by definition, data for the other five countries do not include within-country trade. The figure without intra-euro area trade would be much lower.

7 Summary

This paper studied the transmission of monetary policy with structural time-varying coefficient vector autoregressions in three central European EU members states in comparison with that in the euro area. These countries have experienced changes in monetary policy regimes, which call for the use of a time-varying parameter analysis since. In line with the Lucas Critique, reduced-form models are not invariant to changes in policy regimes. In addition to policy changes, these countries went through substantial structural changes as they became more integrated with the EU. And at a more general level, since the vector autoregression is a linear model, its parameters can change even if the underlying structural model has constant parameters, but it is nonlinear.

Our results indicate that some parameters changed substantially, altering the shape of the impulse response functions. The response of output to a monetary shock has changed in the euro area as well as in the Czech Republic, Hungary and Poland. The time profile of changes are not monotonous, yet in the euro area and also in the three central European economies, monetary transmission became more powerful by 2011 than it had been in 1996. At the last observation of our sample, the fourth quarter of 2011, among the three countries studied, monetary policy was the most powerful in Poland and just slightly less powerful than that in the euro area (though with a different time profile), but the strength of monetary policy was less in Hungary and the Czech Republic. This suggests that, purely from the perspective of the effectiveness of monetary policy, the cost of losing monetary policy independence by entering the euro area would not be large for the Czech Republic and Hungary, while for Poland the loss would be more significant, unless the business cycles of the euro- area and the Polish economy are very closely synchronised.

Due to the small number of countries we studied we could not conduct a formal econometric analysis of the link between the strength of monetary transmission and its underlying determinants, which should be the scope of future research. Yet we assessed the possible role of certain factors, and concluded that the exchange rate regime, the credibility of monetary policy, openness and the share of foreign currency loans may associate well with the differences in the strength of monetary policy.

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9 Appendix: Setting initial conditions

Since the time-varying parameters are assumed to follow random walks, they do not have means that could be used as initial conditions. For this reason we estimated a fixed parameter VAR, equation by equation, with OLS for the first sixteen (effective) observations and used these estimates to form initial conditions for the time varying parameters and their covariance matrix, which is assumed to be block diagonal. For example, when the lag length is 1 and an intercept is not included, the first equation is:

$$y_{1,t} = \phi_{11}^{(1)} y_{1,t-1} + \phi_{12}^{(1)} y_{2,t-1} + \phi_{13}^{(1)} y_{3,t-1} + \phi_{14}^{(1)} y_{4,t-1} + v_{1,t} \,,$$

and let $\hat{\Phi}_1$ denote the estimated coefficient vector, that is, $\hat{\Phi}_1 = \begin{bmatrix} \hat{\phi}_{11}^{(1)} & \hat{\phi}_{12}^{(1)} & \hat{\phi}_{13}^{(1)} & \hat{\phi}_{14}^{(1)} \end{bmatrix}'$. Let $\hat{\Omega}_1$ denote the OLS covariance matrix of this estimation. Let us denote the parameter estimates and their covariance matrices similarly (with different subscripts) for the other three equations.

a_0, the initial condition for the time-varying parameters is set to:

$$a_0 = \begin{bmatrix} \hat{\Phi}_1' & \hat{\Phi}_2' & \hat{\Phi}_3' & \hat{\Phi}_4' \end{bmatrix}'.$$

P_0, the initial condition for the covariance matrix of a_0 is set equal to the block diagonal matrix formed from the OLS covariance matrices of the four equations:

$$P_0 = \begin{bmatrix} \hat{\Omega}_1 & 0 & 0 & 0 \\ 0 & \hat{\Omega}_2 & 0 & 0 \\ 0 & 0 & \hat{\Omega}_3 & 0 \\ 0 & 0 & 0 & \hat{\Omega}_4 \end{bmatrix},$$

where 0 indicates an appropriately sized null-matrix, eg a 4×4 matrix when the lag length is 1 and an intercept is not included.